Review of Socio-Economic Perpectives **Rubino, N. pp. 47-71**
Vol. 3. Issue: 1/ June 2018

DOI: 10.19275/RSEP043
Received: *11.09.2017*
Accepted: *13.03.2018*

EFFECTS OF COMMODITY PRICE LEVELS AND VOLATILITY ON GROWTH IN A LEADING COMMODITY EXPORT FRAMEWORK

Nicola Rubino
Ph.d. candidate, University of Barcelona, faculty of Economics, AQR research group.
E-mail: nicola.rubino.2@gmail.com.

Abstract

Evidence on the commodity curse is nowadays still under debate, with economists focusing on panel data applications in order to uncover the conditional relationship between prices and growth in large sets of countries. While there is much informal evidence to support the 'curse' hypothesis, time series analyses using the VAR methodology have found that increases in commodity prices significantly raise the growth of commodity exporters. In this paper, we adopt cointegration methodology in time series framework for a set of six commodity exporting countries, focusing on the price of their leading exported commodity and attempting to explore the relationship between commodity prices, GDP and growth in a sample covering the 1960-2011 period. After investigating the long run, accounting for a possible break in the series and in the cointegrating relationships, we proceed to analyze the effects of an innovation in GDP in response to price movements through an impulse response function analysis. Our results show evidence of a possible long run relationship between GDP, a set of relevant controls, and the selected commodity prices in three out of the six analyzed countries, conditional on the existence of a single structural break that we loosely identified as the beginning of a transition period to more democratic institutions in each of the analyzed countries. Shocks in the prices of the leading exported commodities we surveyed not only cause a positive response in short run growth, but induce a positive shift in the steady state level of GDP. Such results represent evidence that opposes the idea of the price channel as a medium of transmission of the commodity curse, both in the long and in the short run.

Keywords: Commodity prices, Cointegration, Exports, Growth.

JEL Classification: O11, F31

Citation: Rubino, N. (2018). Effects of Commodity Price Levels and Volatility on Growth in a Leading Commodity Export Framework, Review of Socio-Economic Perspectives, Vol. 3(1), pp. 47-71. DOI: 10.19275/RSEP043.

1. Introduction

A branch of the literature related to economic growth suggests that countries endowed with abundant resources tend to grow at a slower pace than countries where resources are scarce:[1] among the main channels of transmission, theoretical literature has found out three possible channels. First, countries focused on commodity exports face a higher level of volatility than the other countries, due to their lack of specialization. Second, higher prices in the non-tradable economic sectors following the discovery (or the presence) of plentiful natural resources, would lead to the subsequent appreciation of the real exchange rate, causing external losses of competitiveness.[2] Finally, resources which are susceptible to rent seeking might be related to lower quality of the institutions.[3] The existence of a commodity curse, at least in the short run, remains highly debatable. Recent studies have shown how such phenomenon might indeed not constantly hold across resource-dependent countries. In an attempt to verify the first channel of transmission, Raddatz (2005) found out how external shocks, including price shocks, would only account for a small fraction of the volatility of GDP in a vector auto-regression (VAR) application for a selected panel of countries, while, understanding the limitations of the VAR approach and considering the feasibility of a joint short and long run study, most recent mainly focused on dynamic panel data applications in a cointegrating framework,[4] verifying whether or not positive/negative short run effects might be offset in the long run by adverse outcomes. Our study bridges short and long run analysis in a time-series cointegration framework for the 1960-2011 period, studying the effects of prices on growth in six different developing countries. The rest of this paper is organized as follows: in Section 2, a description of the methodology and the data. In Section 3 we report the results of the integration tests for the single series. In Section 4, we report the results of the cointegration tests. In Section 5, the relationships resulting from the past Sections are estimated and discussed. In Section 6, we comment the results of an Impulse response function (IRF) application studying the effects of an oil price shock on growth and long-run GDP. Finally, Section 7 concludes the paper.

[1] See Sachs and Werner (1997), Sachs and Werner (2001) and Sala-i Martin and Subra-manian (2013).
[2] This phenomenon, known as dutch disease, was first highlighted in an article in The Economist in 1977, and later modelled by Corden and Neary (1982) who explained how direct and indirect de-industrialization of less competitive manufacture sectors, following the discovery of new natural reserves, would led to higher price levels in the non-tradable sectors and to less competitive exports, which normally face internationally set prices.
[3] Mehlum, Moene, and Torvik (2006) build on Sachs and Warner works on the resource curse, finding out, contrary to what stated by the latter authors, that abundant resources cause a decrease in aggregate income when institutions are "grabber friendly", while "producer friendly" governments imply a positive relationship between resources and income.
[4] See for instance, Collier and Goderis (2008), Collier and Goderis (2012) and Cavalcanti, Mohaddes, and Raissi (2011), Cavalcanti, Mohaddes, and Raissi (2014). The latter authors in particular, show how both oil abundance would have a positive effect both on income level and economic growth, while volatility of terms of trade would have an overall negative e¤ect, thus disagreeing with Raddatz results in the short term.

2. Methodology and Data

Recent panel data applications have been fundamental in developing a more accurate analysis of the pattern of growth. However, determinants of long run growth across countries have generally been subject to the not always advisable restriction of a common long run behavior in the light of convergence theories.[5] In this paper, we exploited the selection methodology as in Bodart, Candelon, and Carpantier (2012) and Bodart, Candelon, and Carpantier (2015) to select a group of six developing countries based on the average export share of their leading commodity on the total value of exports. Our paper borrows their selection method to capture the relationship between commodity prices and growth by substituting weighted indexes with a commodity price variable for a set of relatively small developing countries. Weighted price indexes show an important drawback: they reflect not just changes in relative prices but also changes in the underlying weights in every country. Furthermore, the relationship be-tween an aggregate index and the analyzed economic variable will depend on the correlation between the prices in the index. This means that the true relationship between growth and prices might be altered by the way prices interact. An index with two perfectly negatively correlated prices would show no relation at all with growth or with any other given variable supposed to depend on it. A way to address the above issues is here presented by exploiting series made of single commodity prices.[6] Another important issue we had to consider was related to the availability of a sufficiently long sample period in yearly frequencies.[7] The majority of the variables we used in the analysis come from version 8.1 of Penn World Table. From this database we took a PPP based real GDP series, called RGDPe, and the total population expressed in millions of

[5] Of course, an heterogeneous estimation in panel data would be feasible, although subject to the degree of heterogeneity across the panel: see Eberhardt and Teal (2010) and Eberhardt and Presbitero (2013) for a few applications to panel estimation accounting for cross sectional correlation and slope heterogeneity based on the mean group estimators family.

[6] Our analysis focused on the idiosyncratic effects on GDP and growth of leading commodity prices movements. As such, an important caveat is that our paper does not currently borrows the convergence-testing strategy framework, based on the stochastic and common trend analysis of GDP seen in Bernanrd and Durlaf (1995) nor from the analysis of the cross sectional variance of the GDP as in Evans (1996). Whenever we refer to changes in the steady state of the economy, we thus refer to the changes of the equilibrium value of long-run GDP in a country by country analysis, without referring to the possibility of the countries examined to be part of a convergence or divergence club. The economies we examine in this paper are transition economies on their path to democratization, which are assumed to be still distant from their equilibrium values. This ultimately explains why, as we will see in the later sections, our estimation strategy had to resort to a subjective selection of the control variables for which we investigated the cointegration property.

[7] A corollary to convergence studies in growth regressions would normally involve some dynamics of output, as Islam (1995) explains in a panel growth context. This leads to the classic formulation of a "Barro" regression with Solowian controls. However, a common characteristic of studies estimating panel growth regressions has been using five or ten years averages to eliminate the possible e¤ect of cyclical fluctuations on long run estimates (see Durlauf, Johnson, and Temple (2005), pag 113, for a brief discussion on the topic). The VECM methodology we employed in this work allowed to analyze both the short and long run effects of commodity prices on growth, and make use of all the possible available information without the need to take averages of the series.

units, pop, to calculate the per capital real GDP; a human capital variable, hc, based on Barro and Lee (2013) hours of schooling variable, and finally a measure of yearly investment and capital accumulation, csh_i. We had to resort to the World Bank's Databank to retrieve a measure of trade openness, tradeop, measured as the ratio between total export plus import volumes over GDP. Finally, the UNCTAD commodities database was used to retrieve the nominal price of Tobacco, Copper, Gold, Coffee, Cocoa, Bananas and Tea, which we deflated by the manufacture unit value index (MUV) in order to construct our series of real commodity prices. All the variables were expressed in logarithms. The countries we considered were selected conditional on the availability of data from the sources we mentioned above and the export weight of their leading exported commodity: due to current limitations in the series at our disposal, we avoided considering oil exporting countries, in particular Nigeria[8] but considered anyway two exporters of mineral commodities, Malawi and Zambia. Finally, out of the fourteen countries suggested by Bodart, Candelon, and Carpantier (2012), six were selected for our analysis: Burundi, Dominica, Ivory Coast, Kenya, Malawi, Mali and Zambia. The implicit long run relationship we thus attempted to estimate, conditional on the outcome of the unit root and the cointegration tests was:

$$RGDP_e = f(hc_t; csh_t; tradeop_t;\ comm_t) \tag{1}$$

Past literature has suggested a large number of possible "controls" in growth regressions. Correlations in the variables we have introduced above and growth has proven to be relatively high,[9] so that, apart from the expected endogeneity issue that would arise from an omitted variable bias, and given the limited amount of periods at our disposal, we choose not to extend this analysis introducing additional control variables. A discussion apart is related to the variables we defined as controls. Cointegration is not a widespread statistical property. Equilibrium variables in the cointegrating vectors might be present or not depending on their integration order, so that their usefulness will vary from country to country, even when eventual breaks in the cointegrating relationship

---

[8] A country which has already been extensively analyzed in Sala-i Martin and Subramanian (2013).

[9] We refer again to the seminal work of Barro (1991) and further developments by Barro and Lee (2013) on the relevance of a negative relationship between growth and initial GDP levels, and a positive relationship between growth and an initial enrollment indicator and overall years of schooling. The use of variable csh_i, the GDP share of gross fixed capital formation, was instead suggested by the work of Levine and Renelt (1992). As Sala-i Martin, Doppelhofer, and Miller (2004) point out after a Bayesian prior selection application, price of investment, primary school enrollment and initial GDP prove to be the most significant and robustly partially correlated variables with long term growth. Finally, the inclusion of the variable tradeop, measuring the degree of trade openness of a country, was motivated by the idea of controlling for the intensity of trade on prices. Even under a dutch disease hypothesis, where the trade open manufacture sector is a price taker, higher trade volumes might lead to a higher price level in the non-tradable sector and subsequent worsening of the resource curse.

are accounted for. This obviously makes the difference, especially in a vector error correction specification, where its main advantage compared to standard VAR analysis is the possibility of observing both long and short run relationships among variables, sometimes at the cost of direct interpretability of the magnitude of coefficients when just a few variables are left after the unit root analysis of the series. In Section 4, we present the results of the residual based and vector based cointegration tests. The residual based unit root tests we employed in the analysis were the Perron and Vogelsang (1992) unit root test with a single structural break for the variables in first differences and the Zivot and Andrews (1992) test for the levels.[10] To account for the issue of small sample bias, we analyzed the series in levels and first differences employing the corrected M-tests from Ng and Perron (2001). As a strategy for unit root testing, we followed the approach suggested by Dickey and Pantula (1987), starting from first differences and then analyzing levels, drawing our conclusion on the order of integration of the series based on the non-rejection of the null hypothesis. After checking for the order of integration of the variables, from which we drew the information we needed to give a structural form to the relationship in equation (1), we will move to Section 4, where we employed alternative specifications of the Gregory and Hansen (1996a) cointegration test to check for the existence of a long run equilibrium among the variables, retrieving in the meanwhile an endogenous break date for each country that we put to use again testing for cointegration in the error correction based approach by Johansen, Mosconi, and Nielsen (2000), based on the Johansen and Juselius (1990) methodologies. We present the results for the cointegrating relationship we obtained in Section 5[11] and discuss an IRF analysis of the relationships for the variables both in their first differences and levels in Section 6.

[10]The Perron and Vogelsang (1992) test was originally specified as a test with a level shift only, while the Zivot and Andrews (1992) allows for a break in the trend and the intercept as well as a break in the intercept only. For the sake of completeness, we present the latter test considering both possibilities. Furthermore, we are aware of the risk of data mining related to the search for a break date in time series. To avoid such critique, we restrict the analysis to testing for a single break only.

[11] The necessity to look for breaks in the variables we presented is essentially due to the changing nature of the economic structures of the economy. Trade openness during the past sixty years has changed substantially across emerging countries, and with it, the path of human capital might have been altered thanks to technological and social spill-over. Moreover, as seen in Christiano (1992) and the time path of GDP in the U.S. after the second world war, visual inspection of the data and pre-estimation selection of the break-date significantly alters the outcome of the analysis.

Table 1: Country selection and export weights

Code	Country	Commodity	Weight
BDI	Burundi	Coffee	51.0%
ICV	Ivory Coast	Cocoa	34.1%
KNY	Kenya	Tea	21.2%
MWI	Malawi	Tobacco	60%
MLI	Mali	Gold	54.1%
ZMB	Zambia	Copper	54.0%

Weights are defined as the ratio between the commodity exports of the country and all the commodity exports of the country. Out of the fourteen countries selected in Bodart, Candelon and Carpantier (2012) we choose those countries whose relative weights were higher than 20%, conditional on the availability of the data.

3. Unit Root Tests

This Section reviews the results on the unit root tests. Results for the Ng and Perron (2001) M-tests are reported in Tables 2 and 3. The additive and outlier models of the Perron and Vogelsang (1992) test for the series' first differences are reported in Tables 4 and 5. The Zivot and Andrews (1992) tests accounting for a break in both the intercept and the trend of the series and the trend only are reported in Tables 6 and 7. According to the Ng and Perron (2001) tests, the majority of the GDPs appear to be integrated of order one, exception made by Burundi's GDP. The same result applied to the Commodity prices, where all the commodities, with the exception of coffee, appear to be I(1). The human capital and the trade opening measure appeared to be I(1) in Burundi and Zambia only, while the gross fixed capital formation variable appear to be non-stationary only in Ivory Coast and Mali. Overall, the outcome of the M-tests did not point at the majority of the variables as suitable long run controls. However, to account for possible break stationarity, we decided to pair-up the M-tests with unit root tests accounting for the presence of one structural break. After the results showed in Tables 4 and 5, based on the first differences, the Zivot and Andrews (1992) tests admitted a much less restrictive set of I(1) variables, once the possibility of a structural break was accounted for. In order to select the variables to carry over to the cointegration tests, since both methodologies, the M-tests and the one break unit root tests, would not suggest a unique solution, we choose to accommodate both suggestions and selected the variable accordingly to both the type of tests we employed, conditional on the fact that the selected variable had been considered I(1) by at least by one of the tests. Our selection can be seen in Table 8.

4. Cointegration Tests

This Section reviews the results on the cointegration tests. Table 10 reports the results for the Gregory and Hansen (1996a) residual-based cointegration tests with one endogenous structural break,[12] Table 11 reports the results for the Johansen cointegration tests, while Table 12 shows the outcome of the Johansen, Mosconi, and Nielsen (2000) cointegration tests with a break in the trend and at a known point in time. We choose to report all model specifications from Gregory and Hansen (1996a), that is model C (a single break in the level of the cointegrating relationship, trend excluded), model C=S=T[13] (a contemporaneous break in the slope of the trend, the slope of the long run elasticity and the intercept), model C=S (a shift in the levels of the relationship and a change in slope of the long run elasticities) and model C=T (a level shift with a trend in the deterministic component). In our test strategy, we choose to compute the results for the Gregory and Hansen (1996a) tests first, so that the endogenous break dates found could be later used for the Johansen, Mosconi, and Nielsen (2000) tests, which require prior knowledge of the known break date. Results from the Gregory and Hansen (1996a) show that an equilibrium relationship could be found in Burundi across all specifications, while little evidence of a complete regime shift in the cointegrating relationship can be found in Mali, where the test for the C=S=T specification could reject the null hypothesis of no cointegration at the 5% level. As explained above, we took the break dates from the residual based tests[14] to compute the Johansen, Mosconi, and Nielsen (2000) tests with some prior knowledge of a possible break date. Results from this latter test, which can be seen in Table 12, confirm the specification for Burundi as being cointegrated, as the test trace statistic reports the presence of at most one cointegrating vector. As for Mali, the weak results in the residual based tests became even weaker evidence in Table 12, where the null hypothesis on no cointegrating relationship was rejected at all conventional levels of significance. For Zambia, non-rejection of the null of no cointegration in Table 10 was followed by non-rejection in the test with a break in Table 12, where the test the presence of at most two cointegrating relationships was detected. Lastly, we observe some weak evidence of cointegration in Ivory Coast in Table 11, which is however not supported by any of the other tests. In conclusion, we choose to estimate two alternative VECM models for Burundi, Zambia and Mali, one with

[12] For the sake of completeness and to compare the results, we also report, in table 9, the results for the Engle and Granger (1987) residual based cointegration tests, which does not account for structural breaks. We included two specifications, one with a constant and another with a constant and a trend. Critical values are taken from MacKinnon (2010).
[13] The C=S=T model was not present in Gregory and Hansen (1996a) but appears in Gregory and Hansen (1996b).
[14] We choose the C=T date for Burundi and the C=S=T break date for Mali, since this two specifications appeared to reject effortlessly the null hypothesis of no cointegration. As for the other countries, our choice fell on the C=T break date.

a specification accounting for the presence of a single break and one without any structural break.

Table 2: Ng and Perron (2001) tests, first differences

	Variable	Mz_a	MZ_t	MSB	MP_t
BDI	RGDP [e]	-0.78	-0.49	0.62	21.65
	hc	-9.57**	-2.09**	0.22*	2.95**
	tradeop	-11.29**	-2.37**	0.21**	2.19**
	comm	-17.98***	-2.94***	0.16***	1.56***
	csh_i	-32.40***	-4.02***	0.12***	0.76***
ICV	RGDP [e]	-27.52***	-3.69***	0.13***	0.92***
	hc	-4.78	-1.55	0.32	5.12
	tradeop	-4.38	-1.38	0.31	5.76
	comm	-12.94**	-2.48**	0.19**	2.14**
	csh_i	-31.93***	-3.99***	0.12***	0.77***
KNY	RGDP [e]	-11.50**	-2.38**	0.21**	2.21**
	hc	-2.16	-0.92	0.43	10.36
	tradeop	-36.36***	-4.23***	0.11***	0.76***
	comm	-2987.64***	-38.65***	0.01***	0.01***
	csh_i	-1.72	-0.80	0.47	12.38
MWI	RGDP [e]	-20.46***	-3.19***	0.16***	1.19***
	hc	-7.04*	-1.81*	0.26*	3.72*
	tradeop	-3.28	-1.27	0.39	7.46
	comm	-10.14**	-2.25**	0.22*	2.42**
	csh_i	-0.01	-0.01	0.39	14.57
MLI	RGDP [e]	-7.25*	-1.87*	0.26*	3.49*
	hc	-11.58**	-2.38**	0.21*	2.21*
	tradeop	-1.06	-0.72	0.68	23
	comm	-7.81*	-1.96*	0.25*	3.18*
	csh_i	-0.56**	-2.29**	0.22**	2.33**
ZMB	RGDP [e]	-7.43*	-1.93*	0.26*	3.31*
	hc	-451.94***	-15.02***	0.03***	0.07***
	tradeop	-56.74***	-5.32***	0.09***	0.44***
	comm	-16.29***	-2.85***	0.17***	1.54***
	csh_i	-3.89	-1.39	0.28	4.45

***Significant at 1%, **significant at 5%, *significant at 1%. The tests include an intercept in the deterministic component.

Table 3: Ng and Perron (2001) tests, levels

	Variable	Mz_a	MZ_t	MSB	MP_t
BDI	RGDP ^{e}	-3.36	-1.29	0.39	27.09
	hc	-12.40	-2.36	0.19	8.03
	tradeop	-8.89	-2.07	0.23	10.38
	comm	-14.53*	-3.42***	0.14***	6.66*
	csh_i	-8.25	-2.01	0.24	11.11
ICV	RGDP^e	-5.08	-1.47	0.29	17.41
	hc	-9.35	-2.14	0.23	9.84
	tradeop	9.41	-2.16	0.23	9.7
	comm	-6.59	-1.77	0.27	13.85
	csh_i	-11.23	-2.34	0.21	8.24
KNY	RGDP^e	-3.64	-1.32	0.36	24.59
	hc	-354020***	-420.73***	0.001***	0.0003***
	tradeop	-15.98*	-2.68*	0.17*	6.56*
	comm	-6.21	-1.58	0.25	14.57
	csh_i	-7.08	-1.80	0.25	12.98
MWI	RGDP^e	-6.68	-1.81	0.27	13.66
	hc	-29.84***	-3.84***	0.13***	3.21***
	tradeop	-7.58	-1.94	0.26	12.04
	comm	-1.85	-0.74	0.4	35
	csh_i	-5.86	-1.71	0.29	15.53
MLI	RGDP^e	-11.65	-2.38	0.2	8
	hc	-26.89***	-3.50***	0.13***	4.34**
	tradeop	-18.41**	-2.98**	0.16**	5.26**
	comm	-8.43	-1.97	0.23	11.06
	csh_i	-9.59	-2.19	0.23	9.49
ZMB	RGDP^e	-4.08	-1.19	0.29	19.82
	hc	-8.22	-1.85	0.23	11060
	tradeop	-2.85	-0.94	0.33	25.19
	comm	-7.88	-1.82	0.23	11.99
	csh_i	-4.23	-2.62	0.19	6.67

***Significant at 1%, **significant at 5%, *significant at 1%. The tests include an intercept in the deterministic component.

Table 4: Perron and Vogelsang (1992) tests, first differences, AO

	Variable	ADF	T1	p-value
BDI	RGDP ^{e}	-7.91**	1975	0.868
	hc	-0.81	1988	0
	tradeop	-5.12**	1994	0.99
	csh_i	-7.87**	1970	0.64
	comm	-6.69**	1999	0.52
ICV	RGDP ^{e}	-6.79**	1978	0.02
	hc	-1.65	1968	0
	tradeop	-6.15**	1972	0.96
	csh_i	-4.62**	1987	0.75
	comm	-1.12	1978	0.33
KNY	RGDP ^{e}	-4.72**	1992	0.29
	hc	-2.23	1992	0
	tradeop	-8.50**	1996	0.48
	csh_i	-1.438	1981	0.97
	comm	-4.84**	1982	0.39
MWI	RGDP ^{e}	-1.25	2003	0.041
	hc_t	-2.93	2008	0.17
	tradeop	-5.03**	1992	0.99
	csh_i	-8.24**	1979	0.56
	comm	-6.85**	1974	0.089
MLI	RGDP ^{e}	-7.28**	1972	0.58
	hc	-2.02	1998	0
	tradeop	-2.89	1972	0.47
	csh_i	-6.15**	1981	0.73
	comm	-0.84	1978	0.64
ZMB	RGDP ^{e}	-6.43**	1996	0.03
	Hc	-1.92	1997	0
	Tradeop	-6.25**	1990	0.58
	csh_i	-2.58	1991	0.21
	comm	-6.67**	2004	0.16

***Significant at 1%, **significant at 5%, *significant at 1%. The tests include an intercept in the deterministic component.

Table 5: Perron and Vogelsang (1992) tests, first differences, IO

	Variable	ADF	T1	p-value
BDI	RGDP [e]	-8.11**	1976	0.06
	hc	-4.92**	1989	0
	Tradeop	-6.58**	1995	0.23
	csh_i	-7.74	1971	0.68
	Comm	-6.65**	1976	0.29
ICV	RGDP [e]	-5.13**	1979	0
	Hc	-1.36	1964	0
	Tradeop	-8.08**	1973	0.78
	csh_i	-7.44**	1988	0.97
	Comm	-7.37	1979	0.13
KNY	RGDP [e]	-7.71**	1993	0.06
	hc_t	-2.48	1989	0.03
	Tradeop	-8.47**	1997	0.19
	csh_i	-1.95	1969	0.47
	Comm	-5.82**	1983	0.23
MWI	RGDP [e]	-6.03**	2004	0.06
	Hc	-3.25	1994	0.12
	Tradeop	-9.09**	1993	0.96
	csh_i	-8.05**	1980	0.98
	Comm	-6.45**	1987	0.06
MLI	RGDP [e]	-7.25**	1973	0.19
	Hc	-3.25	1999	0.01
	Tradeop	-7.55**	1973	0.99
	csh_i	-11.35**	1982	0.72
	Comm	-5.97**	1979	0.19
ZMB	RGDP [e]	-6.53**	1992	0.04
	Hc	-2.89	1994	0.01
	Tradeop	-6.80**	1991	0.15
	csh_i	-2.94	1992	0.17
	Comm	-6.77**	2002	0.06

***Significant at 1%, **significant at 5%, *significant at 1%. The tests include an intercept in the deterministic component.

Table 6: Zivot and Andrews (1992) tests, levels, shift in levels and change in trend

	Variable	ADF	T1
BDI	$RGDP^e$	-3.02	1970
	Hc	-3.97	1981
	Tradeop	-5.97**	1996
	csh_i	-4.85	1989
	Comm	-3.37	1987
ICV	$RGDP^e$	-3.81	1980
	Hc	-4.01	1991
	Tradeop	-5.01	1986
	csh_i	-3.09	1983
	Comm	-3.01	1988
KNY	$RGDPE^e$	-2.71	1977
	Hc	-5.24**	1976
	Tradeop	-4.44**	1997
	csh_i	-5.42**	1982
	Comm	-4.06	1991
MWI	$RGDP^e$	-2.89	1998
	Hc	-3.80	1986
	Tradeop	-7.02**	1981
	csh_i	-4.79	1979
	Comm	-3.86	1969
MLI	$RGDP^e$	-3.83	1985
	Hc	-4.34	1998
	Tradeop	-6.01**	2001
	csh_i	-4.72	1982
	Comm	-2.35	1977
ZMB	$RGDP^e$	-3.06	1992
	Hc	-4.42	1991
	Tradeop	-5.27**	1981
	csh_i	-3.34	1975
	comm	-3.51	1998

***Significant at 1%, **significant at 5%, *significant at 1%. The tests include an intercept in the deterministic component.

Table 7: Zivot and Andrews (1992) tests, levels, change in trend

	Variable	ADF	T1
BDI	$RGDP^e$	-2.77	1978
	hc_t	-4.18	1990
	$tradeop_t$	-2.99	2002
	csh_i_t	-3.49	1982
	$comm_t$	-2.59	2003
ICV	$RGDP^e$	-3.49	1975
	hc_t	-3.79	1982
	$tradeop_t$	-2.95	1992
	csh_i_t	-2.73	1978
	comm	-2.09	1969
KNY	$RGDP^e$	-2.47	1978
	hc	-5.05**	1980
	$tradeop_t$	-4.09	2003
	csh_i	-2.70	1970
	$comm_t$	-4.01	1995
MWI	$RGDP^e$	-2.53	1970
	hc_t	-3.73	1992
	$tradeop_t$	-5.42**	1988
	csh_i_t	-2.75	1995
	$comm_t$	-3.64	1972
MLI	$RGDPE^e$	-3.93	1997
	hc_t	-4.82**	2000
	$tradeop_t$	-6.01**	2002
	csh_i_t	-2.89	1986
	$comm_t$	-2.43	2004
ZMB	$RGDP^e$	-2.87	1997
	hc_t	-4.14	1995
	$Ttradeop_t$	-4.76**	1999
	csh_i_t	-2.38	1985
	$comm_t$	-3.48	2000

***Significant at 1%, **significant at 5%, *significant at 1%. The tests include an intercept in the deterministic component

Table 8: Selected variables after unit root testing variables

BDI	RGDP; hc; tradeop; csh_i
ICV	RGDP; tradeop; csh_i
KNY	RGDP; tradeop
MWI	RGDP; csh_i
MLI	RGDP; tradeop; csh_i
ZMB	RGDP; hc; tradeop

The variables we carried on in the Cointegration analysis selected according to the outcome of the unit root tests.

Table 9: Engle and Granger (1987) cointegration tests

	deterministics	ADF	5% critical value
BDI	trend	-3.53	-5.07
	constant	-2.95	-4.69
ICV	trend	-1.29	-4.34
	constant	-1.16	-4.74
KNY	trend	-2.19	-4.36
	constant	-1.37	-3.91
MWI	trend	-2.49	-4.36
	constant	-2.40	-3.91
MLI	trend	-2.66	-4.77
	constant	-2.41	-4.36
ZMB	trend	-3.11	-4.72
	constant	3.12	-4.32

The tests 5% critical values are taken from MacKinnon (2010)

Table 10: Gregory and Hansen (1996a, 1996b) cointegration tests

	break	ADF	break date	5% critical value
BDI	C=S=T	-7.20**	1993	-6.84
	C=S	-6.17*	1989	-6.41
	C=T	-6.76***	1992	-5.83
	C	-5.66**	1994	-5.56
ICV	C=S=T	-4.72	1978	-6.32
	C=S	-4.12	1993	-6.00
	C=T	-3.69	1970	-5.57
	C	-3.82	1968	-5.28
KNY	C=S=T	-5.22	1995	-5.96
	C=S	-4.73	1996	-5.50
	C=T	-4.42	1998	-5.29
	C	-4.08	1999	-4.92
MWI	C=S=T	-4.81	1988	-5.96
	C=S	-4.43	1992	-5.50
	C=T	-4.61	1992	-5.29

	C	-4.40	1992	-4.92
MLI	C=S=T	-6.22*	1994	-6.32
	C=S	-5.26	1999	-6.00
	C=T	-5.20	2001	-5.57
	C	-4.64	2004	-5.28
ZMB	C=S=T	-4.68	1986	-6.32
	C=S	-4.56	1986	-6.00
	C=T	-4.71	1967	-5.57
	C	-4.43	1986	-5.28

***Significant at 1%, **significant at 5%, *significant at 10%. Maximum number of lags selected based on the and manually imposed, lags based on Schwert (1989) rule of thumb. The number of lags employed is based on a general to specific selection methodology, All the tests include an intercept and a trend.

Table 11: Johansen (1990) cointegration tests

	rank	eigenvalue	trace statistic	5% critical value
BDI	0	-	121.83	68.52
	1	0.68	67	47.21
	2	0.51	33.55**	39.68
ICV	0	-	49.79	47.21
	1	0.47	20.72**	29.68
KNY	0		25.03**	29.68
	1	0.24	12.05	15.41
MWI	0	-	23.39**	29.68
	1	0.23	10.09	15.41
MLI	0	-	47.86	47.21
	1	0.39	23.45**	29.68
ZMB	0	-	56.39	47.21
	1	0.49	22.05**	29.68

**Indicates failure to reject the null hypothesis at 5%. All the tests include an intercept and a trend.

Table 12: Johansen, Mosconi and Nielsen (2001) cointegration tests

	rank	eigenvalue	trace statistic		5% critical value
BDI	0	0.6	123.27	0.65	112.72
	1	0.53	76.93**	0.65	83.5
	2	0.28	39.08	0.65	58.23
ICV	0	0.54	69.35**	0.25	81.32
	1	0.33	32.79	0.25	56.35
	2	0.21	13.85	0.25	35.21
KNY	0	0.39	45.51**	0.75	56.35
	1	0.24	20.91	0.75	35.21
MWI	0	0.31	29.21*	0.65	58.23
	1	0.12	10.59	0.65	36.7

MLI	0	0.55	77.43**	0.75	81.32
	1	0.41	42.98	0.75	56.35
	2	0.29	19.82	0.75	35.21
ZMB	0	0.55	93.52	0.15	77.52
	1	0.46	54.04	0.15	53.31
	2	0.36	22.79**	0.15	32.92

**Indicates failure to reject the null hypothesis at 5%. *Indicates failure to reject the null hypothesis at 10%. All the tests include an intercept and a trend. The tests allow for a break in the trend at a given exogenous point in time.

5. Estimation of the VECM and inference on the long run equilibria

In this Section, we report the estimated vector error correction models for Burundi, Zambia and Mali, for which in Section 4 we found out, conditional on the structure of the deterministic components, at most one cointegrating vector might exist.[15] In equation (2) we report the general specification which we employed to test for cointegration while in equation (3) we report the specification for the cointegration tests with a restricted break in trend.

$$\Delta Y_t = \alpha \begin{pmatrix} \beta \\ \gamma \end{pmatrix}' \begin{pmatrix} Y_{t-1} \\ t \end{pmatrix} + \mu_j + \sum_{i=1}^{k-1} \Gamma_i \Delta Y_{t-1} + \varepsilon_t \quad (2)$$

$$\Delta Y_t = \alpha \begin{pmatrix} \beta \\ \gamma \end{pmatrix}' \begin{pmatrix} Y_{t-1} \\ tD_{t-k} \end{pmatrix} + \mu_j D_{t-k} + \sum_{i=1}^{k-1} \Gamma_i \Delta Y_{t-1} + \sum_{i=0}^{k-1} \sum_{j=2}^{q} \kappa_{j,i} I_{j,t-1} + \varepsilon_t \quad (3)$$

In equation (3) a trend-break dummy variable tD_{t-k} is restricted to the cointegrating vector and represents the trend break in the cointegrating trend-stationary relationship. Y_t represents our vector of variables from Table 8, represents the speed of error correction, is a vector of long run elasticities, is the coe¢ cient of the trend terms, D_{t-k} is the dummy break variable depending on the number of variables $\kappa_{j,i}$ and $I_{j,t}$ is an impulse dummy at the time of the endogenous break. Finally ε_t is a vector of i.i.d. errors. Equation (2) is employed as the basis for our estimation of the linear VECMs, while the break specifications were based on equation (3).[16] In Subsections 5.1, 5.3 and 5.2, we

[15] For the case of Zambia, the Johansen, Mosconi, and Nielsen (2000) tests detected at most two possible cointegrating vectors. However, given the trace statistic rejected the null hypothesis of at most one cointegrating relationship by a very short margin, we choose to conduct the estimation for Zambia considering only one cointegrating relationship.

[16] The use of the impulse break dummies is justified by the need to restrict the residuals to 0 given the initial value in the second Sub-period of our single break analysis. See Joyeux (2007) for further guidance. Note that in both estimations the trend term was restricted to the cointegrating relationship. This allows the cointegrating

will provide a case by case overview of the long run coefficients and loadings, commenting the results, which are visible in the appendix of the article.[17]

5.1. Mali

After its independence in 1960, Mali's democracy started flourishing in 1992 after almost 23 years of military dictatorship. Up until 2011 and after the end of the dictatorship, the country's annual growth rates remained generally positive. It is no surprise that the Gregory and Hansen (1996a) test detected a change in the trend of the equilibrium relationship.[18] Following our test results in the linear VECM, all coefficients of the cointegrating equation appear significant at 5%, exception made for the gold price coefficient, which appears significant only at 10%. The speed of adjustment coefficient indicates a speed of adjustment to the equilibrium values of around 18% each year, three times higher than the average panel 6% estimate reported in Collier and Goderis (2012). The long run elasticity of the real commodity price has a negative sign in the cointegrating vector, capturing a long run positive effect of higher gold prices in Mali. A 10% increase in the price of gold would thus lead to a 1.6% higher long run level of real GDP. Such result is a striking one, albeit well distant from the 5.9% lower GDP level for Nigeria or the overall 16.8% decrease estimated in the panel data analysis of Collier and Goderis (2012). Switching our attention to the estimation accounting for a single break, we observe a strong decrease in the rate of decay of the disequilibrium, which lowers to 2% per year, and an inversion of the sign of the elasticity of GDP with respect to the price of gold, where this time a 10% increase in the price of gold would lead to a 6.2% lower long run level of real GDP. However, the long run relationship appears to be statistically not different from 0, since the loading coefficient is statistically insignificant. The result, on the light of the possibility of a structural break which might have taken place after the beginning of the democratization process in Mali, appears puzzling. Not only the non-existence of the cointegrating vector could be taken as lack of evidence of a possible long run equilibrium, but the lack of clear results after accounting for the democratization process adds up as a potential critique to the rent seeking theory related to the commodity curse or perhaps to the quality of the democratization process in Mali.

relationship to have a trend in equation (2) and a broken trend in equation (3), and the level data to have a trend as well as an unrestricted intercept appears in the deterministic component outside of the relationship.

[17] The complete output is reported in Tables 14 to 18.

[18] We are aware of the fact we might have adopted a set of impulse break dummies to account for years where coups d'état or other events exogenous to the long run equilibrium took place. However, given the limitations imposed by the length of the time series and the eventual loss of degrees of freedom, alongside with the persistent issue of data mining, we choose to focus the analysis on the possibility of an endogenous break only, limiting exogenous components to the deterministic set-up of the estimations.

5.2. Burundi

Since its independence, which took place in 1962, Burundi has been plagued by moments of social tension, which sparked in a series of ethnic conflicts in 1994 between two major ethnic groups, the Tutsu and the Hutu. Only in 2001 a power sharing government, favored by South Africa, allowed for a general cease-fire, granting the first general, democratic election in 2004. However, the authoritarian attitude of the government following the 2010 elections, cast a shadow over the democratic values of the Burundian institutions. In Table 10, the 10 tests suggest the presence of a break in the potential cointegrating relationship somewhere at the beginning of the nineties, result which is furtherly highlighted by the 12 test once 1992 is chosen as a break date. Visibly, any inference on the linear model for Burundi would be biased by the fact that the loading coefficient, which is equal to 0.05 and is not statistically significant. As such, at least in terms of linear modelling, there is no error correction and the cointegration relationship has no economically meaningful interpretation. The nonlinear specification brings forth a similar result, but casts once more doubts over the existence of a long run equilibrium. Remarkably, the loading coefficient would indicate that the economy diverges from an hypothetical equilibrium at the rate of 8% each year. The coefficient for the price of coffee in the cointegrating relationship is negative, implying a positive long run relationship between GDP and coffee prices which would not go against what we would expect from a commodity which is not subject to rent capture, but it is statistically not different from 0.

5.3. Zambia

Contrary to the countries analyzed up until now, since its peaceful independence in 1964, Zambia has been characterized by a relatively stable government and overall political stability. The Gregory and Hansen (1996a) test indicates a potentially meaningful break date in the cointegrating relationship in 1986, a date perhaps distant from the late nineties, when the privatization of the mining sector, mainly based on copper extraction, began to draw the attention of foreign investors, increasing FDI and output. Results in the linear specification could be immediately interpreted as confirming the existence of a long run resource curse effect regardless of the degree of political stability in the country. The result appears to side with the historical vision of Sachs and Werner (1997) rather than with those authors who support the thesis of the political channel as a mean of transmission of the resource curse. The estimated long run elasticity of -1.87 is found to be in the range of the elasticities discovered by Collier and Goderis (2012) for the non agricultural commodity export price index, but leads to an excessive reaction of long run GDP levels to changes in the Commodity price. The relationship appears to be non-interpretable, since the loading coefficient appears statistically not significant. Moving to the nonlinear specification, the loading coefficient for the first difference of Real GDP, significant at the 1% level, appears to indicate an even faster half-life of disequilibrium at a convergence rate of 27% per year. Once again, the convergence rate appears very high relative to our benchmark index estimation, Collier and Goderis (2012), while in the cointegrating relationship no evidence of a negative effect of mineral prices over GDP in the long run could be found, given the negative sign of the elasticity of GDP to Copper prices and the statistical significance of the coefficient.

6. VECM and VAR impulse response function analysis

This Section switches the analysis from the static inference on the cointegrating vector to the computation of the transitory or permanent effect of an innovation in prices to the long run equilibrium level of GDP the economies we studied appear to converge to. We present, based on the past VECMs, some evidence that a sudden price variation at a given point in time might permanently affect the long run equilibrium level of the economies in the countries were we found our system of variables to be cointegrated.[19] We set up an impulse response analysis based on variable levels form of the countries were we found the cointegration relationship to exist and be statistically significant.[20] This results have been paired up with results from orthogonalized impulse responses based on a VAR system with an innovation equal to one standard deviation, which we employ to test for short/medium run effects of price shocks to growth. The ordering strategy of the variables for the Cholesky decomposition was organized such that a one standard deviation impulse in the innovation of the international price variables has a contemporaneous impact on all the other variables following along the lower triangular matrix, as we would normally expect in commodity-dependent countries with no price setting power, and consistently with the Dutch disease assumptions. The ordering of the residual variables would then be tradeop, hc, csh_i and finally RGDP. In other terms, this ordering implies that a shock in the degree of trade openness in a small open country is very unlikely to have a contemporaneous effect on the price of the most exported commodity, but will have strong effects on the remaining variables because of the state of dependence of the economy exports on that variable, affecting human capital and physical capital accumulation. Similarly, we expect a shock on human capital to have almost immediate effects on investment decisions, but

[19] The cointegrating vectors we identified in the past section should be seen as "static" relationships, where any alterations of the equilibrium set by the elasticities is corrected at a given speed according to the loading coefficient. This section instead, investigates the commodity curse effect in terms of responses to shocks, both in the levels and the first difference of the variables. Evidently, conclusions drawn in this section will differ from those on the equilibrium relationship in the past section.

[20] Although a Wold moving average representation, required for impulse response function analysis, an impulse response matrix can be computed for a VECM as if it was a VAR. So, IRFs are available even if some or all the variable are not I(0). Impulses hitting a nonstationary system, can have a permanent effect on the shocked variables. We would expect the results of the visual analysis of the error forecast functions to be not dissimilar from the conclusions we draw from the inference on the cointegrating vector in the previous section.

no contemporaneous effects on the other variables since the degree of the trade openness variable partially depends on imports, which we assume are exogenous. Finally, we expect a shock to the variable related to gross fixed capital formation, csh_i, to feed into GDP immediately but not directly into the remaining variables. Percentile confidence intervals for the VECM and the orthogonalized VAR were bootstrapped with one hundred drawings with replacement from the centered residuals of the VAR and the IRF calculated for a twenty years period.

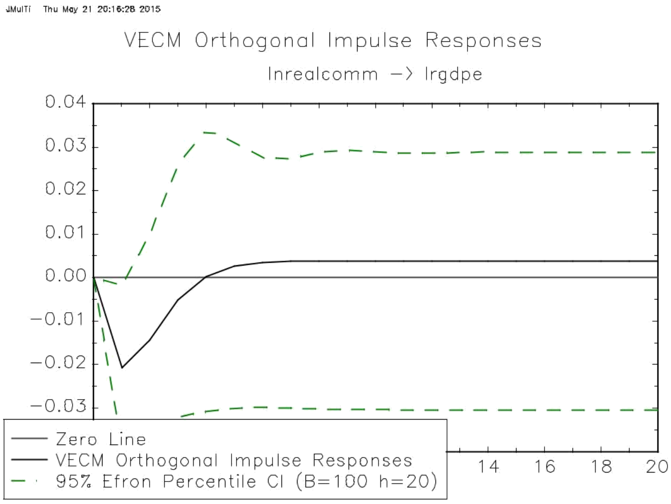

Figure 1: Mali, Levels IRF

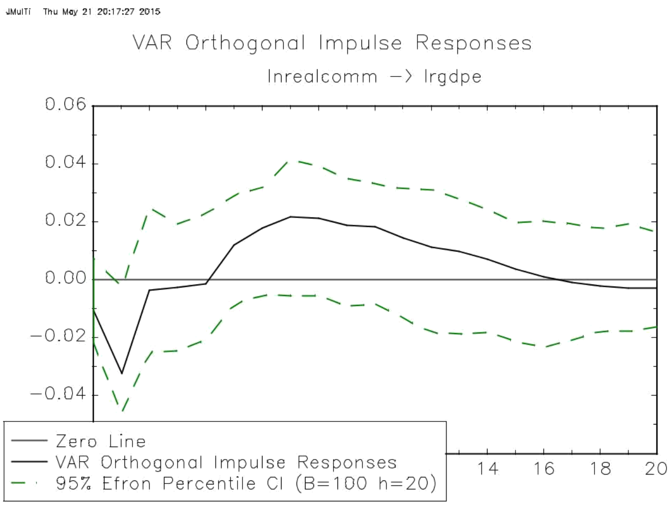

Figure 2: Mali, first differences IRF

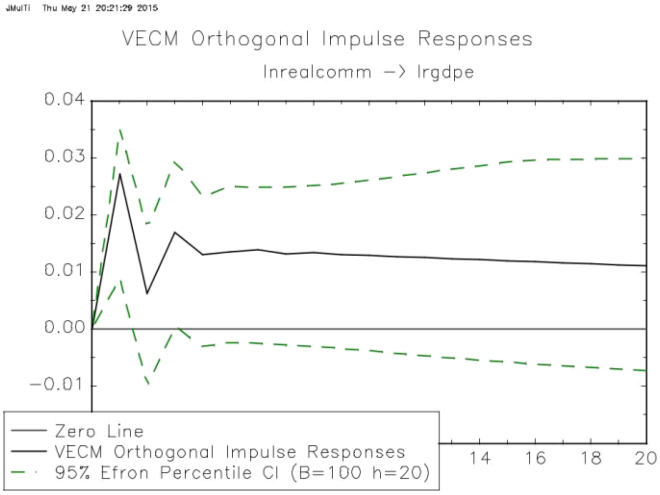

Figure 3: Burundi, Levels IRF

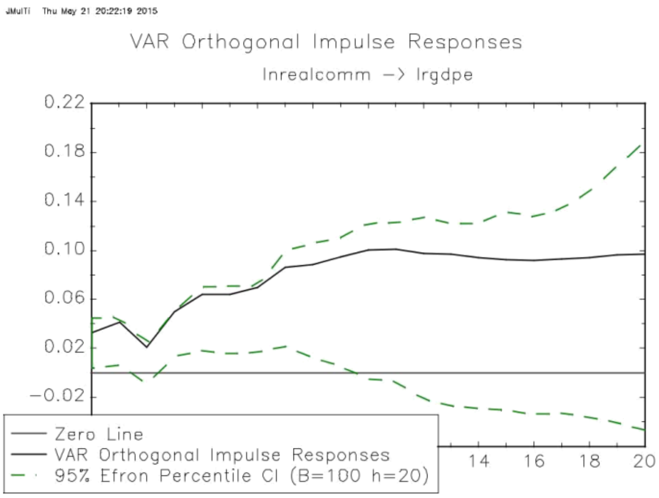

Figure 4: Burundi, First Differences IRF

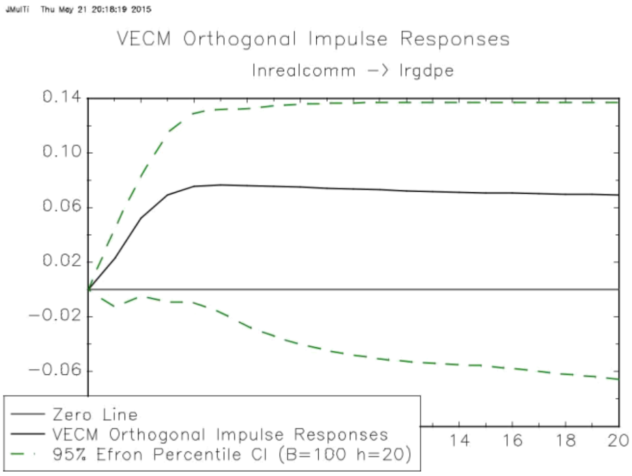

Figure 5: Zambia, Levels IRF

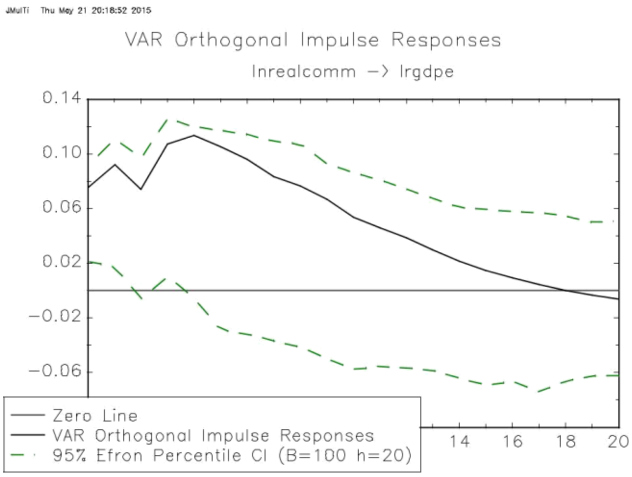

Figure 6: Zambia, First Differences IRF

The analysis of the IRF Figures 2, 3 and 5 would then allow us to check whether or not the effect of a shock to the commodity prices causes a permanent change in the long run equilibrium levels, while Figures 1, 4 and 6 can be interpreted as the 20 years horizon transitory effect of price deviations on GDP growth.[21] As we have already pointed out, the results are interpreted on the light of the price transmission channel of the resource curse hypothesis affecting the long run steady state GDP and growth in the short/medium run. In Mali, a one standard deviation shock to the price of gold appears to have no substantial effect on the long run GDP level (Figure 2). The shock however, appears to have a short and negative five periods effect on growth in the country (Figure 1), with an initial negative jump at time 0, after which the response variable becomes positive for around thirteen periods and then converges again to its initial state. In Burundi, an innovation in the price of coffee appears to generate a permanent effect on the GDP equilibrium level (Figure 3). The shock in the first differences VAR also appears to cause a persistent short run positive response on growth, which lasts for at least as long as the horizon of the IRF (Figure 4). In Zambia, a positive shock to the price of Copper implies a positive shift of the equilibrium level of GDP in the long run (Figure 5). GDP growth response to the price of copper remains consistently positive, with the effect dying out only eighteen periods after the initial shock (Figure 6). Overall, economic growth appears to be positively influenced by positive variations in the commodity prices, with the exception of an initial, negative response in Mali to a shock in the gold price, and long run GDP to be positively affected by price shocks, indicating no evidence of the commodity curse price channel hypothesis in the short or the long run.

7. Conclusions

Our paper analyzed the effect of changes in the price of leading exported commodities in a group of commodity exporting countries. Our results shows evidence of a possible long run relationship between GDP, a set of relevant controls, and the selected commodity prices in three out of the six analyzed countries, conditional on the existence of a single structural break that we loosely identified as the beginning of a transition period to more democratic institutions in each of the analyzed countries. Our estimations led us to conclude that shocks in the prices of leading exported commodities not only cause a positive response in short run growth, but influence as well the steady state level of GDP in the long run. Finally, such results represent evidence rejecting the idea of the price channel as a medium of transmission of the commodity curse, both in the long and in the short run.

[21] Note that we did not make any comments on the bootstrapped confidence intervals, as one hundred drawings might perhaps not be enough to safely make inference based on them.

References

Barro, R. J. (1991): *"Economic Growth in a Cross Section of Countries"*, Quarterly Journal of Economics, 106(2), 407–443.

Barro, R. J., and J. W. Lee (2013): *"A new dataset of educational attainment in the world, 1950-2010"*, Journal of Development Economics, 104, 184–198.

Bernanrd, A. B., and S. N. Durlaf (1995): *"Convergence in International Output"*, Journal of Applied Econometrics, 10, 97–108.

Bodart, V., B. Candelon, and J. F. Carpantier (2012): *"Real exchange rates in commodity producing countries: a reapprisal"*, Journal of international money and finance, 31(6), 1482–1502.

Bodart, V., B. Candelon, and J. F. Carpantier (2015): *"Real exchange rates, commodity prices and structural factors in developing countries"*, Journal of International Money and Finance, 51(C), 264–284.

Cavalcanti, T. V. d. V., K. Mohaddes, and M. Raissi (2011): *"Growth, development and natural resources: New evidence using a heterogeneous panel analysis"*, The Quarterly Review of Economics and Finance, 51, 305– 318.

Cavalcanti, T. V. d. V., K. Mohaddes, and M. Raissi (2014): *"Commodity Price Volatility and the Sources of Growth"*, Journal of Applied Econometrics, DOI: 10.1002/jae.2407.

Christiano, L. J. (1992): *"Searching for a break in GNP"*, Journal of Business & Economic Statistics, 10(3), 237–249.

Collier, P., and B. Goderis (2008): *"Commodity Prices and Growth: Reconciling a Conundrum*", Munich Personal RePEc archive, (17315).

Collier, P., and B. Goderis (2012): *"Commodity prices and growth: an empirical investigation"*, European Economic Review, 56, 1241–1260.

Corden, W. M., and J. P. Neary (1982): *"Booming Sector and De-Industrialisation in a Small Open Economy"*, Economic Journal, 92(368), 825–848.

Dickey, D. A., and S. G. Pantula (1987): *"Determining the Order of Differencing in Autoregressive Processes"*, Journal of Business & Economic Statistics, 5(4), 455–461.

Durlauf, S. N., P. A. Johnson, and J. R. W. Temple (2005): *"Growth Econometrics",* Handbook of Economic Growth, 1(8), 555–667.

Eberhardt, M., and A. F. Presbitero (2013): *"This Time they are Different: Heterogeneity and Nonlinearity in the Relationship between Debt and Growth"*, IMF Working Paper, 248.

Eberhardt, M., and F. Teal (2010): *"Productivity analysis in global manufacturing production"*, Department of Economics, Oxford discussion papers.

Engle, R., and C. W. J. Granger (1987): *"Co-integration and Error Correction: Representation, Estimation and Testing"*, Econometrica, 55, 251–276.

Evans, P. (1996): *"Using cross-country variances to evaluate growth theories"*, Journal of Economic Dynamics and Control, 20(6-7), 1027–1049.

Gregory, A. W., and B. E. Hansen (1996a): "*Residual-based tests for Cointegration in Models with Regime Shifts"*, Journal of Econometrics, 70, 99–126.

Gregory, A. W., and B. E. Hansen (1996b): *"Test for Cointegration in Models with Regime and Trend shifts", Oxford Bulletin of Economics and Statistics*, 58(3), 555–560.

Islam, N. (1995): *"Growth Empirics, A Panel Data Approach"*, The Quarterly Journal of Economics, 110(4), 1127–1170.

Johansen, S., and K. Juselius (1990): *"Maximum Likelihood Estimation and Inference on Cointegration - With Applications to the Demand for Money"*, Oxford Bulletin of Economics and Statistics, 52, 169–210.

Johansen, S., R. Mosconi, and B. Nielsen (2000): *"Cointegration analysis in the presence of structural breaks in the deterministic trend"*, Econometrics Journal, 3, 216–249.

Joyeux, R. (2007): *"How to deal with structural breaks in practical cointegration analysis"*, Cointegration for the Applied Economists, pp. 195–222.

Levine, R., and D. Renelt (1992): *"A Sensitivity Analysis of Cross-Country Growth Regressions"*, The American Economic Review, 82(4), 942–963.

MacKinnon, J. G. (2010): *"Critical Values for cointegration Tests"*, Queen's Economics Department Working Paper No. 1227, Queen's University.

Mehlum, H., K. Moene, and R. Torvik (2006): *"Institutions and the re-source curse"*, The Economic Journal, 116, 1–20.

Ng, S., and P. Perron (2001): *"Lag length selection and the construction of unit root tests with good size and power"*, Econometrica, 69(6), 1519–1554.

Perron, P., and T. Vogelsang (1992): "*Nonstationarity hypothesis and level shifts with an application to purchasing power parity"*, Journal of Business and Economic Statistics, 10, 301–320.

Raddatz, C. (2005): "*Are External Shocks Responsible for the Volatility of output in Low - Income Countries?*", World Bank Policy Research Working Paper, 3680.

Sachs, J. D., and A. M. Werner (1997): "*Sources of Slow Growth in African Economies*", Journal of African Economies, 6(3), 335–376.

Sachs, J. D., and A. M. Werner (2001): "*The curse of natural resources*", European Economic Review, 45, 827–838.

Sala-i Martin, X., G. Doppelhofer, and R. I. Miller (2004): "*Determinants of Long-Term Growth: a Bayesian Averaging of Classic Estimates (BACE) approach*", The American Economic Review, 94(4), 813–835.

Sala-i Martin, X., and A. Subramanian (2013): "*Addressing the Natural Resource Curse: an Illustration from Nigeria*", Journal of African Economies, 22(4), 570–615.

Zivot, E., and W. K. Andrews (1992): "*Further Evidence on the Great Crash, the Oil-Price Shock and the Unit-Root Hypothesis*", Journal of Business & Economic Statistics, 10(3), 251–270.